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Abstract

We study wage effects of two important elements of non-wage labour costs: firing costs and payroll taxes. We exploit a reform that introduced substantial reduction in these two provisions for unemployed workers aged less than thirty and over forty five years. Theoretical insights are gained with a matching model with heterogeneous workers, which predict a positive effect on wages for new entrant workers but an ambiguous effect for incumbent workers. Difference-in-differences estimates, which account for the endogeneity of the treatment status, are consistent with our model predictions and suggest that decreased firing costs and payroll taxes have a positive effect on wages of new entrants. We find larger effects for older than for younger workers and for men than for women. Calibration and simulation of the model corroborate such positive effect for new entrants and also show a positive wage effect for incumbents. The reduction in firing costs accounts, on average, for one third of the overall wage increase.

Keywords: Dismissal costs, payroll tax, evaluation of labour market reforms, difference-in-difference, matching model, Spain.

JEL classification: C23, D31, J31.

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1 Introduction

In the last decades, several European countries have reduced employment protection and payroll taxes to improve the performance of the labour market (see Kugler (2007) for employment protection legislation (EPL) reforms and Carone, Nicodme, and Schmidt (2007) for recent changes in payroll taxes).[1] However, the estimation and evaluation of the causal effects of the changes has proved difficult, since most changes have been gradual (i.e. not sharp) and accross the board (i.e. applied to everyone).

In 1997, Spain drastically reduced dismissal costs and payroll taxes for young and old workers only, which provides a unique natural setting to examine the effects of non-wage labour costs. Severance payments for unfair dismissals were reduced 20%, while payroll taxes decreased between 40% and 60%, depending on the targeted group. These sharp changes, which applied only to some age groups, provide a unique opportunity to examine the causal effects of firing costs and payroll taxes on employment and wages.

There is an increasing amount of empirical evidence, which points that stringent employment protection regulations reduce employment flows (Autor, Donohue, and Schwab (2004, 2006); Kugler and Pica (2003, 2008)). However, evidence on wage effects is very scarce and not very conclusive. Leonardi and Pica (2010) analyse an increase in firing costs implemented in Italy for small firms and find that more stringent employment protection has a negative impact on entry and subsequent wages, while van der Wiel (2010) finds positive wage effects of extending employer's term of notice in the Netherlands.

The incidence of payroll taxes also gathers mixed evidence. Generally speaking, when employees percieve a close link between employers' contributions and their benefits, payroll taxes are likely to be fully shifted from firms to employees, with no disemployment effects. However, with a loose link between taxes and benefits, payroll taxes are usually not fully passed on to employees and employment decreases.[2] Small changes have also been found easier to pass on to employees than large changes (Gruber (1997)).

[1]For instance, in the late 1980s France relaxed employment protection provisions to facilitate employment for certain types of workers, and Germany has recently (in 2004) exempted small firms (from 5 to 10 employees) from EPL. Payroll taxes decreased in the EU-27 from 7.5% to 7.3% of GDP between 1995 and 2005, and the Nordic countries have been reducing payroll taxes selectively for some regions since the mid 1980s.

[2]This may be the case for pay-as-you-go social security systems, such as the Spanish one, with weak linkages between pensions and other benefits, on the one hand, and contributions, on the other.

Our analysis focusses on the wage effects of firing costs and payroll taxes.[3] To do so, we extent the matching model with heterogeneous workers put forth by Dolado, Jansen, and Jimeno (2007) in two important ways to accomodate the salient features of the Reform. We consider the joint effect of payroll taxes and firing costs on wages, and since the reform basically targets the entry wage of two groups of workers, we consider a different wage bargaining process for new entrants than for incumbent workers. The theoretical model predicts a positive effect on wages for new entrant workers but an ambiguous effect for incumbent workers. While the effect of payroll taxes is always negative, firing costs increase the wage for incumbents and reduce the wage of new entrants. This result takes place because firing costs are only operational for incumbents but not for new hired workers, therefore the bargaining power of the formers is relatively higher.

We provide two sets of complimentary evidence, from estimations and from simulations, which yield consistent results. Estimates come from a microeconometric analysis of panel individual administrative records, while simulations are obtained by first calibrating the model and then simulating the reform.

We exploit the variation of firing costs and payroll taxes across age groups (young, prime-age, and older) and over time (before and after 1997), and identify the effects of the reform using a difference-in-differences estimator, i.e. we compare wages of younger and older individuals with those of prime-age individuals, before and after the reform.

Our main findings suggest that decreased firing costs and payroll taxes have a positive effect on the wages (and employment) of new hired workers. Estimated effects are larger for older than for younger workers and for men than for women. Our simulations show that decreased firing costs and payroll taxes also have a positive effect on the wages of incumbent workers, and that on average payroll taxes account for over two thirds of the overall wage increase. Such relative contribution is lower for new entrants than for incumbent workers.

The experience of Spain should also provide direct evidence on the effects other countries might expect from a decision to promote (permanent) employment by reducing non-wage labour costs. Since firing costs and payroll taxes account for a large proportion of overall non-wage labour costs in many countries, they are likely to be used in the future

[3] In Cervini Plá, Ramos, and Silva (2010), the companion and more extensive working paper, we also consider the implications of firing costs and payroll taxes on employment.

to boost employment, as they have been extensively used in the past. Our results suggest that a substantial cut in non-wage labour costs has an important and substantial effect.

Our paper contributes to the small but growing literature that uses large policy changes within a country over time or across groups to evaluate their labour market effects. Our analysis makes several advances over previous studies. Unlike many previous studies we present a taylor-made model that fits the salient features of the policy changes. On the empirical side, we provide new evidence on the wage effects of non-wage labour costs. The data we use is a unique longitudinal data set which contains information on individual job histories from social security records and basic individual information from the census. Thus, we can work with all relevant job spells instead of quarterly data, as provided for instance by the Labour Force Survey. We use information on previous unemployment spells to overcome the sample selection problem we face when estimating the causal effects on wages, which results from those getting new employment not being a random sample.

The rest of the paper is structured as follows. Next we briefly describe the main changes brought about by the 1997 Spanish labour market reform, while Section 3 accommodates the salient features of the reform into a matching model with heterogenous workers. Section 4 explains our identification strategy and section 5 presents the data. Our main estimation results are reported in Section 6. Finally, section 7, summarises the main findings of the paper.

2 Institutional background

Employment protection legislation and especially firing costs have undergone substantial changes in the last twenty five years in Spain. In the early 1990s, nearly one third of overall employment in Spain was temporary –twice the European average–, and nearly all new hires signed temporary contracts (Guell and Petrongolo (2007)), which entailed lower severance payments than permanent contracts when separation took place earlier than agreed or nil when the termination date was observed, and whose termination could not be appealed. Such a rapid increase in temporary employment, brought about by a liberalisation in the use of temporary contracts that took place in 1984, led to a dual labour market (insider-outsider) and segmentation problems between unstable low-paying jobs and stable high-paying jobs (Dolado, García-Serrano, and Jimeno (2002)).

In order to increase the share of permanent employment, and after a first unsuccessful reform in 1994,[4] the 1997 reform substantially lowered firing costs for unfair dismissals and payroll taxes to newly signed permanent contracts, when the worker belonged to certain population groups. In particular, severance payments for unfair dismissals were cut by about 25% and payroll taxes fell between 40% and 90% for new permanent contracts of workers younger than 30 years old, over 45 years old, the long-term unemployed, long-term unemployed women who enter to under-represented occupations, and disabled workers. We only exploit the differential treatment by age group, since the long-term unemployed and women under-represented in their occupations may be self-selected, and disabled workers are a very distinct group which deserves a separate analysis. In particular, we study newly signed permanent contracts from unemployment. Conversions of temporary to permanent contracts after the second quarter of 1997 were also promoted with reductions in dismissal costs and payroll taxes for some population groups —see Appendix Table 8. However, since the reductions were very similar across age groups, identification of the effects becomes less clear-cut and therefore we will not use this group either. Table 1 shows the principal changes in key provisions introduced by the 1997 reform for the younger and older workers. Severance payment for targeted groups were reduced from 45 to 33 days' wages per year of seniority and the maximum time period was reduced by half, from 24 to 12 months. Reductions in payroll tax differ by age group; they fall by 60% and 40% for older and younger unemployed individuals, respectively for a period of 24 months. After the first 24 months, a lower payroll tax reduction of 50% is extended indefinitely only for individuals over 45 years of age.

Social security contribution rebates decreased slightly for newly signed contracts in 1999 and these changes were eventually extended in 2001.[5] These further changes in

[4]The new regulations introduced with the 1994 reform restricted the use of temporary contracts to seasonal jobs and tried to reduce dismissal costs for permanent contracts by relaxing the conditions for 'fair' dismissals of workers under permanent contracts. In particular, the definition of 'fair' dismissal was widened by including additional 'economic reasons' for dismissals. However, as Dolado, García-Serrano, and Jimeno (2002) point out, in practice, not much changed: employers continued to hire workers under temporary contracts for all type of jobs —and not only for seasonal jobs—, and judges did not change their behaviour when appraising dismissals, despite the new regulations, i.e. dismissals under 'economic reasons' continued to be granted mainly when there was agreement between employers and workers, so labour courts continued to rule most dismissals as unfair.

[5]In particular, payroll taxes were reduced 35% in the first year and 25% in the second year for newly hired young unemployed workers under permanent contract, while reductions for older unemployed workers were 45% for the first year and 40% for the second one. Dismissal costs, however, did not change in 1999.

Table 1: **Principal Changes in Dismissal Cost and Payroll Tax due to the Labour Market Reform of 1997 which permit identification for Unemployed Workers**

		Dismissal cost under existing permanent contracts (pre-reform)	Dismissal cost under new permanent contracts (post-reform)	Payroll tax reductions for newly hired workers under permanent contracts after 1997
Treated groups	Young (<30 years)	45 days' wages per year of seniority with a maximum of 42 months' wages	33 days' wages per year of seniority with a maximum of 24 months' wages	40% of employer contribution for 24 months
	Older (>45 years)	45 days' wages per year of seniority with a maximum of 42 months' wages	33 days' wages per year of seniority with a maximum of 24 months' wages	60% of employer contribution for 24 months, 50% thereafter
Control group	Middle-aged (30-45 years)	45 days' wages per year of seniority with a maximum of 42 months' wages	45 days' wages per year of seniority with a maximum of 42 months' wages	None

provisions, though minor, will condition our sample period to one year before and after the reform, i.e. 1996 and 1998 (see Section 4).

3 A theoretical framework

In order to analyze the wage effects of the 1997 reform, this section uses the matching model with heterogeneous workers put forth by Dolado, Jansen, and Jimeno (2007) with two extensions. First, we illustrate the joint effects of payroll taxes and firing costs on wages. Second, since the reform basically targets the entry wage of two groups of workers (less than 30 years and more than 45 years old, respectively), we consider a different wage bargaining process for new entrants than for incumbents workers. This distinction is relevant because the firm does not incur in firing costs when the firm and the worker do not agree on a wage in the first encounter since a contract has not yet been signed. This second assumption permits deliver theoretical predictions specific to the entry wage, which is the dependent variable of the micro estimates.

This labour market consists of a measure 1 of risk-neutral, infinitely-lived workers and a continuum of risk-neutral, infinitely-lived firms. Workers and firms discount future payoffs at a common rate δ and capital markets are perfect. In addition, time is discrete.

There are three type of workers, young (y), middle-age (m) and elderly (e) workers who can be either unemployed or employed. The employed can be either new entrants or incumbents. Thus, there are six type of workers who earn w_{0t}^j and w_t^j, where subscript 0 indicates new entrants and superscript $j = y, m, e$ denotes the age-group of workers. There is a time-consuming and costly process of meeting unemployed workers and job vacancies. As in den Haan, Ramey, and Watson (2000), we assume that the meeting function takes the following form

$$M(u_t, v_t) = \frac{u_t v_t}{(u_t^\varphi + v_t^\varphi)^{1/\varphi}}, \qquad \varphi > 0, \tag{1}$$

where u_t denotes the unemployment rate and v_t are vacancies. This constant-return-to-scale matching function ensures that ratios $M(u_t, v_t)/u_t$ and $M(u_t, v_t)/v_t$ lie between 0 and 1. Due to the CRS assumption they only depend on the vacancy-unemployment ratio θ_t. The former represents the probability at which unemployed workers meet jobs, $f(\theta_t) = M(1, 1/\theta_t)$. Similarly, the latter denotes the probability at which vacancies meet workers, $q(\theta_t) = M(\theta_t, 1)$. Each period, there is a proportion $\lambda_t^j = u_t^j/u_t$ of each type of workers looking for jobs.

Firms have a production technology that uses only labour. Each firm consists of only one type of job which is either filled or vacant. Before a position is filled, the firm has to open a job vacancy with cost c per period. A firm's output depends on aggregate worker's productivity A_t^j and a match-specific term z_t. The match-specific productivity term z_t is assumed to be independent and identically distributed across firms and time, with a cumulative distribution function $G(z)$ and support $[0, \bar{z}]$. We assume there is a productivity gap between each type of worker.

Every period, a proportion ϕ^j of each type of employed worker separate exogenously from the employment status and flow into the unemployment pool. Firms may voluntarily terminate employment relationships, for which they may incur in a firing cost. In particular, firms lose γ^j when a match with a incumbent worker is destroyed by the firm. In this case, a proportion ψ of this cost is assumed to be a transfer to the worker in form of severance payment whereas the rest $(1 - \psi)$ is assumed to be fully wasted, reflecting firing restrictions imposed by the government. These costs are not operational during the job meeting process. The second policy parameter is the wage payroll tax to be paid by the firm, τ^j.

The equations characterizing the value of vacancies, V_t, and filled positions for new jobs, $J^j_{0t}(z_t)$ and incumbent jobs $J^j_t(z_t)$ are, [6]

$$
\begin{aligned}
V_t &= -c + \lambda^y_t \delta \left[q(\theta_t) \int_{\tilde{z}^y_{0t+1}}^{\bar{z}} J^y_{0t+1}(z) dG(z) + [1 - q(\theta_t)(1 - G(\tilde{z}^y_{0t+1}))] V_{t+1} \right] & (2)\\
&+ \lambda^m_t \delta \left[q(\theta_t) \int_{\tilde{z}^m_{0t+1}}^{\bar{z}} J^m_{0t+1}(z) dG(z) + [1 - q(\theta_t)(1 - G(\tilde{z}^m_{0t+1}))] V_{t+1} \right] \\
&+ (1 - \lambda^y_t - \lambda^m_t) \delta \left[q(\theta_t) \int_{\tilde{z}^e_{t+1}}^{\bar{z}} J^e_{0t+1}(z) dG(z) + [1 - q(\theta_t)(1 - G(\tilde{z}^e_{0t+1}))] V_{t+1} \right], & (3)\\
J^j_{0t}(z_t) &= A^j_t z_t - (1 + \tau^j) w^j_{0t}(z_t) + \delta(1 - \phi^j) \left[\int_{\tilde{z}^j_{t+1}}^{\bar{z}} J^j_{t+1}(z) dG(z) + G(\tilde{z}^j_{t+1}) \left(V_{t+1} - \gamma^j \right) \right] \\
&\quad + \delta \phi^j V_{t+1}, & (4)\\
J^j_t(z_t) &= A^j_t z_t - (1 + \tau^j) w^j_t(z_t) + \delta(1 - \phi^j) \left[\int_{\tilde{z}^j_{t+1}}^{\bar{z}} J^j_{t+1}(z) dG(z) + G(\tilde{z}^j_{t+1}) \left(V_{t+1} - \gamma^j \right) \right] \\
&\quad + \delta \phi^j V_{t+1}, & (5)
\end{aligned}
$$

where $\tilde{z}^j_{0t+1}$ and $\tilde{z}^j_{t+1}$, $j = \{y, m, e\}$, are match-specific productivity thresholds, defined such that nonprofitable matches (i.e., with negative surplus) are severed. These thresholds or reservation productivities must satisfy the following conditions:

$$
\begin{aligned}
J^j_{0t}(\tilde{z}^j_{0t}) - V_t &= 0, & (6)\\
J^j_t(\tilde{z}^j_t) - V_t + \gamma^j &= 0. & (7)
\end{aligned}
$$

Expression (6) defines the reservation productivity associated to the hiring process of unemployed workers who meet a vacant job. Note that in this case the firm is not entailed to γ in the absence of agreement since the job has not been created yet. In turn, (7) defines the reservation productivity for job destruction of existing positions. In this case, firing costs γ become operational.

It follows that each type of worker separate and find jobs with probabilities,

[6] For exposition reasons, we omit writing the aggregate state variables $\{A_t, \theta_t\}$ as arguments of these value functions.

$$s_t^j = \phi^j + (1-\phi^j)G(\widetilde{z}_t^j), \tag{8}$$

$$\chi_t^j = f(\theta_{t-1})(1 - G(\widetilde{z}_{0t}^j)). \tag{9}$$

$$\tag{10}$$

On the workers' side, each type of unemployed worker gets b^j units of the consumption good each period, which could be understood as the value of leisure, home production, or unemployment benefit. The values of the different statuses - unemployed, U_t^j, new hired $W_{0t}^j(z_t)$ or incumbent, $W_t^j(z_t)$ - are given by the following expressions:

$$\begin{aligned}
U_t^j &= b^j + \delta\left[f(\theta_t)\int_{\widetilde{z}_{0t+1}^j}^{\bar{z}} W_{0t+1}^j(z)dG(z) + [1 - f(\theta_t)(1 - G(\widetilde{z}_{0t+1}^j))]U_{t+1}\right], \quad (11)\\
W_{0t}^j(z_t) &= w_{0t}^j(z_t) + \delta\left[(1-\phi^j)\left(\int_{\widetilde{z}_{t+1}^j}^{\bar{z}} W_{t+1}^j(z)dG(z) + G(\widetilde{z}_{t+1}^j)\left(U_{t+1}^j + \psi\gamma^j\right)\right)\right]\\
&+ \delta\phi^j U_{t+1}^j, \quad (12)\\
W_t^j(z_t) &= w_t^j(z_t) + \delta\left[(1-\phi^j)\left(\int_{\widetilde{z}_{t+1}^j}^{\bar{z}} W_{t+1}^j(z)dG(z) + G(\widetilde{z}_{t+1}^j)\left(U_{t+1}^j + \psi\gamma^j\right)\right)\right]\\
&+ \delta\phi^j U_{t+1}^j. \quad (13)
\end{aligned}$$

To close the model, we need first to incorporate two additional assumptions. One is the free entry condition for vacancies: firms will open vacancies until the expected value of doing so becomes zero. Therefore, in equilibrium we must have

$$V_t = 0. \tag{14}$$

The other assumption is that wages are set through Nash bargaining. The Nash solution is the wage that maximizes the weighted product of the worker's and firm's net return from the job match. The first-order conditions for new and incumbent employees yield the following conditions,

$$\begin{aligned}
(1-\beta)(1+\tau^j)(W_{0t}^j(z_t) - U_t^j) &= \beta(J_{0t}^j(z_t) - V_t), \quad (15)\\
(1-\beta)(1+\tau^j)(W_t^j(z_t) - U_t^j - \psi\gamma^j) &= \beta(J_t^j(z_t) - V_t + \gamma^j). \quad (16)
\end{aligned}$$

Note that the Nash condition for the incumbents displays two extra terms depending on γ. Since separation costs are now operational, they are explicitly considered in the wage negotiation. This implies that the firm's threat point when negotiating with an incumbent is no longer the value of a vacancy V_t, but $(V-\gamma)$; and that the worker's threat point depends on the proportion of firing costs (ψ) obtained in case of disagreement. Defining the total surplus for new and incumbent jobs as,

$$S_{0t}^j(z_t) = (1+\tau^j)(W_{0t}^j(z_t)-U_{0t}^j)+(J_{0t}^j(z_t)-V_t), \tag{17}$$

$$S_t^j(z_t) = (1+\tau^j)(W_t^j(z_t)-U_t^j-\psi\gamma^j)+(J_t^j(z_t)-V_t+\gamma^j), \tag{18}$$

and using (3)-(16), the equilibrium wage equation for new entrants and incumbents are

$$\begin{aligned} w_{0t}^j(z_t) &= (1-\beta)b^j-\left(1-\phi^j\right)\delta\left(1-\beta\right)\psi\gamma^j-\frac{\beta}{(1+\tau^j)}\left(1-\phi^j\right)\delta\gamma^j \\ &+\frac{\beta}{(1+\tau^j)}\left[A_t^j z_t+\delta f(\theta_t)(1-\beta)\int_{\tilde{z}_{0t+1}^j}^{\bar{z}} S_{0t+1}^j(z)dG(z)\right], \end{aligned} \tag{19}$$

$$\begin{aligned} w_t^j(z_t) &= (1-\beta)b^j+\left[1-\left(1-\phi^j\right)\right]\delta\left(1-\beta\right)\psi\gamma^j+\frac{\beta}{(1+\tau^j)}\left[1-\left(1-\phi^j\right)\right]\delta\gamma^j \\ &+\frac{\beta}{(1+\tau^j)}\left[A_t^j z_t+\delta f(\theta_t)(1-\beta)\int_{\tilde{z}_{0t+1}^j}^{\bar{z}} S_{0t+1}^j(z)dG(z)\right], \end{aligned} \tag{20}$$

where

$$\begin{aligned} S_{0t}^j(z_t) &= A_t^j z_t-(1+\tau^j)(1-\beta)b^j-\delta\left(1-\phi^j\right)\left[1-\left(1+\tau^j\right)\psi\right]\gamma^j \\ &-\delta f(\theta_t)\beta\int_{\tilde{z}_{0t+1}^j}^{\bar{z}} S_{0t+1}^j(z)dG(z)+(1-\phi^j)\delta\int_{\tilde{z}_{t+1}^j}^{\bar{z}} S_{t+1}^j(z)dG(z), \end{aligned} \tag{21}$$

It is immediate to see that direct effects of firing costs γ and pay roll taxes τ go in the same direction in entry wages, w_{0t}^j, but in the opposite direction in continuing wages, w_t^j. In both cases, payroll taxes decrease the wages of a new and continuing workers because they reduce the net share of the match product obtained by the worker. In turn, firing costs decrease w_{0t}^j because these costs are not operational at the entry level jobs, reducing the workers 'implicit' bargaining power. Notice that the higher the proportion of severance payments in total firing costs ψ, the larger the negative effect of firing costs on w_{0t}^j. Thus, the overall effect of the 1997 reform on the wages of workers who made the transition from unemployment to jobs with permanent contracts should be positive. In the case of w_t^j, however, job firing costs increase w_t^j because they become operational, increasing the workers 'implicit' bargaining power. Now the higher the proportion of severance payments in total firing costs ψ, the larger the positive effect of firing costs on w_t^j. Thus, the overall effect of the 1997 reform on 'average wages' of workers with permanent contracts is entirely an empirical question.

To fully characterize the dynamics of this economy, we need to define the law of motion for unemployment and the mass of employed workers (u_t^j and n_t^j). These evolve according to the following difference equations:

$$
\begin{aligned}
n_t^j &= n_{t-1}^j + f(\theta_{t-1})\lambda_t^j(1 - G(\widetilde{z}_{0t}^j))u_{t-1}^j - s_t^j n_{t-1}^j && (22)\\
n_t &= n_t^y + n_t^m + n_t^e && (23)\\
u_t^j &= u_{t-1}^j + s_t^j n_{t-1}^j - f(\theta_{t-1})\lambda_t^j(1 - G(\widetilde{z}_{0t}^j))u_{t-1}^j, && (24)\\
u_t &= u_t^y + u_t^m + u_t^e, && (25)\\
1 &= u_t + n_t, && (26)
\end{aligned}
$$

4 Identification strategy

In order to identify the impact of dismissal costs and payroll taxes on wages, we compare the change in mean wages of young and older employees holding a permanent contract in the current spell and who were unemployed in the previous spell before and after the 1997 reform, with the change in mean wages of middle age workers who got a permanent job from unemployment. That is, we exploit the variation over time and across age groups and use a difference-in-differences estimator. The identifying assumption requires that the difference between wages of treatment and control groups would not change in the absence of the reform. More formally,

$$
E\{\tilde{w}_{pre}^T\} - E\{\tilde{w}_{pre}^C\} = E\{\tilde{w}_{post}^T\} - E\{\tilde{w}_{post}^C\}
$$

where $\tilde{w}$ is the counterfactual wage in absence of the reform, superscript $j = T, C$ indicates treatment or control group and subscripts *pre* and *post* refer to pre- and post-reform periods.

In the empirical analysis, we identify the average effect of the reform on wages as:

$$\beta_{DID} = (E\{w^T_{post}\} - E\{w^T_{pre}\}) - (E\{w^C_{post}\} - E\{w^C_{pre}\}) \tag{27}$$

where w is actual wages. The identification strategy is illustrated in Figure 1, which plots average wages for men and women by age group relative to the second quarter of 1997, for the years before and after the reform, i.e. 1995 to 1999. Figure 1 shows a marked change in the growth rate of average wages of the treatment groups, after the reform. That is, after the second quarter of 1997 average wages of younger and older workers increase much faster than those of the control group, and the increase is larger for men and for the older age group.

As treatment and control groups consist of individuals who make a transition from unemployment to permanent employment, they are likely not to be a random sample since some individual characteristics may determine the probability of entering permanent employment from unemployment. We take account of this sample selection problem with a two-step Heckman type correction, and identify the first step (i.e. the probability of making a transition to permanent employment from unemployment) with two variables that characterize the unemployment history of the individual: number of unemployment spells prior to the transition and unemployment duration over all spells.

We estimate the effect of the reform on wages with the following wage equation:

$$W_{it} = \alpha_0 + \alpha_1 D_t + \alpha_2 D_i + \beta^{'} D_i \times D_t + X^{'}\gamma + \delta\lambda + \epsilon_{it} \tag{28}$$

where W_{it} is the log of gross monthly earnings for those who transit from unemployment to permanent contract, D_i is a vector of dummies for treated groups (i.e. workers who make a transition to permanent employment from unemployment and are aged less than 30 or older than 45 years) and D_t is a vector of dummies that identify the post-reform years. The vector X includes time-varying covariates such as education, occupation and industry. The coefficients of interest in this regression are the βs, which represent the treatment effects; that is, capture the effects of the reform on wages in the years after the reform. Finally, λ is the selection coefficient, which derives from the following first stage linear probability model:

$$Pr[e_{it} = 1|X_{it}] = \Lambda[\mu_o + \mu_1 D_t + \mu_2 D_i + \theta^{'} D_i \times D_t + X^{'}\gamma + Z^{'}\eta] \tag{29}$$

where $e_{it} = 1$ if individual i transits from unemployment to permanent employment and $e_{it} = 0$ otherwise. The vector Z includes the two variables that help identify this first step regression, that is, the number of unemployment spells prior to the transition and unemployment duration over all spells.

Our strategy assumes that employers do not substitute workers not affected by the reform for targeted workers. However, if the change in provisions brought about by the reform is perceived as beneficial by employers, they will tend to substitute non-targeted workers (our control group) for targeted workers (our treatment group) who are otherwise deemed similar. To see whether the assmumption holds, Table 2 presents pre- and post-reform employment probabilities for individuals with ages adjacent to the relevant age thresholds, i.e. 30 and 45 years. If employers substituted workers, pre- and post-reform employment probabilities for control group workers would fall. Table 2 shows that employment probabilities for these workers do not change significantly, which suggests that the possible substitution of workers is not likely to affect our results. To further check whether substitution is a problem we estimated the effects on employment of the reform with the sample restricted to the narrower defined age treatment and control groups. If substitution took place then we would find larger effects in the restricted sample. Results of these regressions, presented in Cervini Plá, Ramos, and Silva (2010), show that this is not the case: the effects of the reform on employment probabilities estimated with the restricted sample are quite similar to (and usually slightly smaller than) the effects obtained with the whole sample.[7]

5 Data and methodological decisions

We employ a unique administrative dataset with Social Security records called Countinuous Sample of Job Histories (Muestra Continua de Vidas Laborales, MCVL) for the year 2005, which consist of a random sample of 4% of all affiliated workers, working or not, and pensioners from the Social Security archives. This dataset contains detailed job-related information on the complete job history of of 1,142,118 individuals, which include labour market status and type of contract for each and every job spell.[8] The MCVL is very rich

[7]The only exception is the unemployment to permanent employment transition probability of older women.

[8]Since the dataset contains information also on pensioners, we do *not* face attrition problems due, for instance, to the larger likelihood of workers with poorer employment performance and lower wages to

Table 2: **Pre- and Post-reform employment probabilities for a restricted sample**

	Men		Women	
Age	Pre-reform	Post-reform	Pre-reform	Post-reform
27	50.2%	56.0%	45.2%	49.1%
28	53.1%	59.0%	46.7%	50.1%
29	55.0%	62.0%	48.7%	52.1%
30	60.3%	60.5%	51.2%	52.3%
31	62.8%	63.0%	55.9%	55.7%
32	64.0%	63.9%	56.1%	57.2%
42	70.1%	70.7%	65.1%	66.1%
43	71.3%	71.8%	66.3%	66.8%
44	71.2%	71.1%	67.4%	68.0%
45	73.5%	76.3%	67.9%	70.3%
46	73.0%	76.9%	69.3%	74.2%
47	74.1%	79.1%	70.1%	73.6%

and detailed as regards job histories, but lacks information on basic individual characteristics. To this end, we match the MCVL and municipal information (padrones) and recover individual information on sex, education and age.

Our sample selection is as follows. First, we study men and women aged between 21-60 to select out the two ends of the labour career. Second, we drop the long-term unemployed and disabled workers. The former may be a self-selected group while the latter are a very distinct group which deserves a separate analysis. Third, we only use job spells posterior to 1993, since prior to that year information on type of contract is not reliable. Fourth, we drop incomplete or incorrect registers. Fifth, we consider workers who are in the general scheme (Regimen General), which includes 90 per cent of all workers; i.e. we exclude the self-employed, workers in Agriculture, Fishing and other minor special cases.[9] To avoid capturing the effects of the 1999 reform, we compare the year prior to the reform (1996) with the year after the reform (1998). Sensibility checks are performed with slightly wider time windows (i.e.1995-1996 and 1998-1999), but results do not change substantially (see Appendix Table 9).

The wage measure is the log of gross monthly wage or salary, deflated by the consumer

exit the labour market sooner.

[9]This is common practice in the few studies that use the MCVL (e.g. García-Pérez and Rebollo-Sanz (2009)) and is also the choice of Kugler, Jimeno, and Hernanz (2002) when studying the employment effects of the reform using the Spanish Labour Force Survey.

price index. As it often occurs with Social Security records, wages in the MCVL are top- and bottom-coded, that is, they are censored. Although for the entire sample this is a significant problem (Bonhomme and Hospido (2009)), such an issue is likely not to be empirically relevant in our case as wages are censored only for very few observations.[10]

Tables 3 and 4 provide summary statistics by relevant age groups of our sample for men and women separately. Descriptive statistics are presented for the period before and after the 1997 Reform. The last three rows suggest that the probability of getting a permanent contract or to make a transition from temporary to permanent employment might have increased after the reform and especially so for the youngest age group.

Table 3: Descriptive Statistics by Age Group, Pre- and Post-Reform for **Men**

Variable	Age<30		Age 30-45		Age>45	
	Pre	Post	Pre	Post	Pre	Post
Wages	1017.33	1138.80	1308.47	1397.11	1399.34	1548.18
Log wages	6.925	7.048	7.177	7.242	7.240	7.345
Age	25.15	24.96	36.57	36.15	51.94	52.25
% Incomplete Primary Education	12.80	16.07	21.59	25.84	45.18	50.12
% Primary Education	43.47	45.68	35.50	39.42	28.24	28.12
% Secondary and Technical Education	37.19	32.89	35.55	28.65	20.47	17.09
% University	6.54	5.37	7.36	5.99	6.01	4.68
% with Permanent Contract	43.89	53.28	75.54	76.69	82.07	83,81
% with Temporary Contract	56.11	46.67	24.46	23.28	18.93	16.19
Unemployment spells	4.54	4.66	3.24	3.67	2.35	2.75
Unemployment duration	1017.63	897.64	1123.21	1075.13	1409.93	1228.51
N	26,443	70,394	49,950	102,144	26,973	50,867

The matched MCVL has important advantages over other data sets which have been employed in previous studies. For instance, as compared with the Spanish Labour Force Survey (Encuesta de Población Activa, EPA), used by Kugler, Jimeno, and Hernanz (2002) to examine the effects of the reform on employment, the MCVL contains information on wages for each job spell, which allows us to examine the effects on wages, for

[10]There are hardly any bottom-coded observations in our sample, while top-coded wages represent between 0.16% and 0.66% of the sample, depending on the year and sex group. Such small incidence is likely to be due to the fact that individuals in our sample have experience a recent spell of unemployment, so their wages are less likely to be affected by top-coding.

Table 4: Descriptive Statistics by Age Group, Pre- and Post-Reform for **Women**

Variable	Age<30		Age 30-45		Age>45	
	Pre	Post	Pre	Post	Pre	Post
Wages	987.55	1045.78	1305.56	1399.76	1377.76	1401.76
Log wages	6.90	6.95	7.17	7.24	7.23	7.25
Age	24.94	24.93	36.28	35.97	51.76	51.70
% Incomplete Primary Education	7.98	8.22	19.3	18.10	49.52	48.85
% Primary Education	33.57	33.98	35.15	35.62	28.22	31.06
% Secondary and Technical Education	47.80	47.04	35.66	36.06	17.16	15.55
% University	10.65	10.76	9.89	10.21	5.1	4.55
% with Permanent Contract	40.38	41.81	67.35	70.49	70.79	73,52
% with Temporary Contract	59.62	58.19	32.65	29.51	29.21	26.48
Unemployment spells	4.90	4.97	3.83	4.08	2.90	3.32
Unemployment duration	1071.63	942.17	1224.94	1163.55	1517.24	1362.98
N	30,886	118,854	34,969	103,510	10,940	28,602

first time. Secondly, the MCVL provides information on each and every single job spell and not only at the time of the interview, as typically occurs with other large and representative surveys such as the European Community Household Panel (ECHP), the EU Survey on Income and Living Conditions (EU-SILC), or the Labour Force Surveys, which eliminates the possibility of aggregation bias. The time-span of the MCVL, however, is not long enough as to cover more than one economic cycle, and thus cycle effects cannot be taken account of in the empirical analysis.

6 Wage Effects of the 1997 Reform

As pointed out in the Introduction, we present two sets of complementary evidence on the effects of the 1997 reform. We first present microeconometric estimates (Section 6.1) and then evidence which results from calibrating and simulating the model of Section 3 (Section 6.2). Difference-in-differences estimates will yield reults for men and women separately, while results from simulations provide effects on average wages across gender. Simulations, however, permit compute the effect on wages for new entrants and incumbent workers. Finally, simulations will also allow us to calculate the separate effect of dismissal costs and payroll taxes.

6.1 Microeconometric estimates

Table 5 reports the estimates of interest of the wage equation (28) in the upper panel and of the selection equation (29) in the lower panel, for men and women separately. The effect of the reform on wages is captured by the coefficients β on the interaction $(D_i \times D_t)$, which is positive and statistically significant for the two treatment groups and both genders. This means that the reduction in dismissal costs and payroll taxes results in a sizeable wage increase for the two treated groups as compared to the control group. The increase is larger for the older group than for the younger one and smaller for women than for men. More precisely, we find a 3.5% wage increase for young unemployed men transiting to a permanent contract; the increase for women of the same age is lower (2.7%). For the older unemployed workers doing the same transition, wages increased 7.6% and 5.4% for men and women, respectively. [11]

Table 5: Effects of the Reform on Wages for men and women who experience a transition from unemployment to permanent employment

	Men		Women	
	Coefficient	t-stat	Coefficient	t-stat
Wage equation				
Age<30	-0.264	-25.49	-0.189	-15.75
Age>45	0.060	2.71	0.054	2.51
(Age<30)*(Post 1997)	0.035	5.49	0.027	4.51
(Age>45)*(Post 1997)	0.076	5.28	0.054	3.35
Selection coeff (λ)	17.469	43.55	12.434	29.66
Selection equation				
Age<30	-0.065	-6.21	-0.104	-12.31
Age>45	-0.062	-6.34	0.035	3.45
(Age<30)*Reform	0.025	4.26	0.019	3.88
(Age>45)*Reform	0.085	14.64	0.035	9.36
Unemployment spells	-0.033	-33.54	-0.008	-14.54
Duration	-0.0002	-21.56	-0.0001	-32.51

Notes: Control group are men and women aged 30 to 45 years.

Controls have the expected sign. For instance, age dummies (D_i) show a monotonic

[11] Recall that long-term unemployed and disabled workers were dropped from the sample because all individuals belonging to these two groups receive treatment irrespective of their age.

and positive relationship between age and wages. Full estimates of the wage and selection regressions are shown in Appendix Table 10.

Selection into the relevant transition from unemployment to permanent employment is indeed not random, but positive, i.e. unobservables are positively correlated with both doing the transition and wages, as indicated by the positive and statistically significant δ. The two coefficients of the variables that identify selection into the relevant transition η are negative and statistically significant. That is, a larger number of unemployment spells or longer overal time in unemployment reduces the probability of signing a permanent contract from unemployment.

6.2 Calibration and simulated results of the model

In this section we quantify the impact in relative wages when only new hired workers of each target group are assumed to be directly affected by the reform. To this end, we first calibrate the model presented in section 3 at annual frequencies just before the 1997 labour market reform. Then we departure from the initial setup by reproducing the observed reduction in firing costs and payroll tax in the targeted age-groups during the 1997 reform. Finally, we analyze the simulated post reform effects on the level of wages of each target group with respect to the non targeted group of workers (m). The simulated results complement the estimated effects presented in section 6.1 by predicting not only the impact on wages of newly hired workers, w_0^j, but also on wages of continuing workers not directly affected by the reform, w^j.

6.2.1 Benchmark calibration: Before the reform

Our benchmark parametrization must match the following targets in the steady state, which are summarized in the upper part of Table 6. The first three targets consist of the average unemployment rates for workers younger than 30 years old, $u^y = 34.7\%$, between 30 and 45 years old, $u^m = 18.0\%$, and older than 45 years old, $u^e = 12.3\%$. Using data from the Spanish National Institute of Statistics (INE), we apply Shimer (2005)'s methodology to target an annual job finding rate of 0.555 for young workers, 0.441 for middle age employees and 0.450 for older employees. We also target the average wage differential among these groups. Thus, $\overline{w}^y/\overline{w}^m = 0.777$ and $\overline{w}^e/\overline{w}^m = 1.069$. In the absence of data on hiring costs for Spain, we rely on Abowd and Kramarz (2003) and target hiring costs equivalent to 3.2% of annual labour costs per worker in France, which has a level of employment protection similar to the Spanish one.

With respect to the calibration of our parameters, we set the discount factor $\delta = 0.95$, which matches an annual real interest rate of nearly 5 percent observed in 1996. The workers' bargaining power β is set to 0.5. Petrongolo and Pissarides (2001) identify an elasticity of unemployment with respect to the matching function in the range 0.5-0.7. We take 0.6 as reference and thus set the matching parameter φ at 0.869.

Using data from the OECD Tax Database, we set the payroll tax at 0.30 for all groups. Thus, $\tau^y = \tau^m = \tau^e = 0.30$. Next we turn to the firing costs γ^j. We first estimate the total severance payments in years of wages for permanent contracts, $\psi\gamma^j$, using the following information from Osuna (2005): (i) 20 days of wages per year of seniority for legal indemnities in fair dismissals with a maximum of 12 monthly wages; (ii) 45 days of wages per year of seniority for unfair dismissals with a maximum of 42 monthly wages dismissals; (iii) the mean job tenure X^j for each worker-age group; (iv) procedural wages of around two monthly wages; and (v) the fact that 72% of all firing processes were declared unfair in 1996.

According to our target unemployment and job finding rates, the calibrated job exit rates of each group are $s_t^y = 0.294$, $s_t^m = 0.097$ and $s_t^e = 0.065$. These rates imply that the average job tenure in 1996 was around 3 years for employees younger than 30 years, 10.3 years for those workers between 30 and 45 years, and 15.4 years for employees older than 45 years old. Thus, severance payments amount to $\psi\gamma^y = 0.518 \times w^y$, $\psi\gamma^m = 1.242 \times w^m$ and $\psi\gamma^e = 1.769 \times w^e$ of annual wages.[12]

We next calculate the firing tax costs, $(1 - \psi)\gamma^j$. Garibaldi and Violante (2005) estimate it between 19% and 34% of total firing costs, depending on the layoff scenario. We consider the last scenario and set ψ equal to 0.66. Thus, the firing tax component amounts to near 51.5% of severance payments, which implies that $(1-\psi)\gamma^y = 0.267 \times w^y$, $(1-\psi)\gamma^m = 0.640 \times w^m$ and $(1-\psi)\gamma^e = 0.911 \times w^e$.[13] Finally, total firing costs are equal to $\gamma^y = 0.785 \times w^y$, $\gamma^m = 1.882 \times w^m$ and $\gamma^e = 2.680 \times w^e$.

[12]For X^j years of job tenure severance payments in years are $\psi\gamma^j$=(0.72$\times X^j \times$45 days per year + 0.28$\times X^j \times$20 days per year + 60 days)/365.

[13]The annual firing tax calculation amounts to $(1 - \psi)\gamma^j = \psi\gamma^j \times 0.515 \times w^j$.

Table 6: **Benchmark Calibration. Spain, 1996**

		Value	Source
Targets:			
Unemployment rate (< than 30 years old)	u^y	0.347	[A]
Unemployment rate between 30 and 45 years old	u^m	0.180	[A]
Unemployment rate > than 45 years old	u^e	0.123	[A]
Job finding rate > than 45 years old	χ^e	0.450	[A]
Job finding rate between 30 and 45 years old	χ^m	0.440	[A]
Job finding rate < than 30 years old	χ^y	0.550	[A]
Wage gap for young workers	$\frac{\overline{w}^y}{\overline{w}^m}$	0.777	[A]
Wage gap for old workers	$\frac{\overline{w}^e}{\overline{w}^m}$	1.069	[A]
Hiring costs	$\frac{c}{\overline{w}}$	0.032	[B]
Parameters:			
Aggregate labour productivity > than 45 years old	A^e	1.000	Normalized
Aggregate labour productivity between 30 and 45	A^m	0.991	[C]
Aggregate labour productivity < than 30 years old	A^y	0.850	[C]
Mean of log z	μ	0.000	Normalized
Standard deviation of log z	σ_z	0.10	[D]
Discount rate	δ	0.950	[A]
Exogenous exit probability > than 45 years old	ϕ^e	0.061	[C]
Exogenous exit probability between 30 and 45	ϕ^m	0.083	[C]
Exogenous exit probability < than 30 years old	ϕ^y	0.294	[C]
Employment opportunity cost < than 30 years old	b^y	0.558	[C]
Employment opportunity cost between 30 and 45	b^m	0.753	[C]
Employment opportunity cost > than 45 years old	b^e	0.763	[C]
Employers payroll tax	τ^j	0.300	[B]
Cost of vacancy	c	0.027	[C]
Parameter of the Matching function	φ	0.869	[D]
Worker's bargaining power	β	0.50	[D]
Total firing costs parameter < than 30 years old	γ^y	$0.785w^y$	[A,B]
Total firing costs parameter between 30 and 45	γ^m	$1.882w^m$	[A,B]
Total firing costs parameter > than 45 years old	γ^e	$2.680w^e$	[A,B]
Proportion of severance payments	ψ	0.66	[B]

Note: [A] Own calculation based on original data; [B] Other studies;
[C] Obtained from model to match the targets; [D] Own assumption

Following the standard assumption in the literature, as in den Haan, Ramey, and Watson (2000), the idiosyncratic productivity z_t is assumed to be log-normally distributed with mean μ and standard deviations σ_z. We normalize the mean of $\log z_t$ to zero, $\mu = 0$. With respect to σ_z, and similar to den Haan, Ramey, and Watson (2000), we set it equal to 0.1. We also normalized the aggregate labour productivity for the group of workers with more than 45 years old, $A^e = 1.00$, and fix $A^y = 0.850$ and $A^m = 0.991$ to match the observed wage gap among these workers. Finally, the hiring cost c is calibrated together with the employment opportunity costs b^j and with the exogenous job exit probability ϕ^j. We select these parameters to satisfy the hiring costs target of 3.2% of average wages, as well as our remaining calibration targets: $u^y = 34.7\%$, $u^m = 18.0\%$, $u^e = 12.3\%$, $\chi^y = 0.550$, $\chi^m = 0.440$ and $\chi^e = 0.450$. This yields $c = 0.027$, $b^y = 0.558$, $b^m = 0.753$, $b^e = 0.764$, $\phi^y = 0.294$, $\phi^m = 0.083$ and $\phi^e = 0.061$.

6.2.2 Simulated effects

The first principal change in legislation reduced severance payments by around 20% for workers who made the transition from unemployment to permanent jobs (33 days of wages per year of seniority, with a maximum of 24 monthly wages, rather than 45 days of wages per year of seniority with a maximum of 42 monthly wages in case of unfair dismissal).[14] The second main modification of the reform was a reduction of 40% and 60% in the payroll tax for workers under 30 and over 45 years of age who made the transition from unemployment to permanent jobs. Thus, for new hired workers, τ^y and τ^e are reduced from 0.30 to 0.18 and 0.12, while it remains unchanged at 0.30 for both the middle aged group and the continuing positions of the young and elderly groups. As in the empirical part, the simulation takes into account the changes experienced by wages just after the reform. The results of this exercise are displayed in the first panel of Table 7.

The simulated reform yields a similar increase in the relative wage of the two target groups. With respect to the group of new hired workers older than 45 who made the transition from unemployment to jobs with permanent contracts, the simulated ratio w_0^e/w_0^m increased by 7.14%. This result is in line with the estimated effects reported in

[14] In this case, for new hired workers with age-group $i = y, e$, the calculations are: $\psi\gamma^i$=(0.72×XX^i years×*33 days* per year + 0.28×XX^i years×20 days per year + 60 days)/365. Thus, the annual firing tax calculation amounts to $(1-\psi)\gamma^i = \psi\gamma^i \times 0.515 \times w^i$.

Table 7: **Simulated effects of the 1997 reform**

Simulated post reform variation		Var.(%)
New hired worker wages ratio: w_0^y/w_0^m		6.91
New hired worker wages ratio: w_0^e/w_0^m		7.14
Incumbent worker wages ratio: w^y/w^m		1.29
Incumbent worker wages ratio: w^e/w^m		3.16
Average wages ratio: $\overline{w}^y/\overline{w}^m$		3.21
Average wages ratio: $\overline{w}^e/\overline{w}^m$		4.24
Unemployment rate (%):	Pre-reform	Post-reform
u^y	34.7	33.3
u^i	18.0	18.2
u^o	12.6	9.8

Estimated effects of the 1997 reform.	
Weighted average* first panel of Table 5	Var.(%)
New hired worker wages ratio: w_0^y/w_0^m	3.19
New hired worker wages ratio: w_0^e/w_0^m	6.75

Simulated post reform variation with no reduction in γ	Var.(%)
New hired worker wages ratio: w_0^y/w_0^m	2.68
New hired worker wages ratio: w_0^e/w_0^m	3.25
Incumbent worker wages ratio: w^y/w^m	1.16
Incumbent worker wages ratio: w^e/w^m	2.28
Average wages ratio: $\overline{w}^y/\overline{w}^m$	2.21
Average wages ratio: $\overline{w}^e/\overline{w}^m$	2.68

*Note: Weighted average of the estimates of table 5, where weights are population shares of the two gender groups.

section 6.1. For instance, the estimated wage effect for old unemployed workers who do a transition to permanent contract is a wage increase of 6.75% relative to their middle-aged counterpart, as it can be seen in the second panel of Table 7. In turn, the relative wage for younger unemployed workers who do the same transition increases by 6.91%, which is larger than the estimated one of 3.19%.

Notice that in spite of the absence of adjustment in the firing costs and payroll taxes of continuing workers with permanent contracts, the model simulates an increase of 1.29% and 3.16% in the ratios w^y/w^m and w^e/w^m, respectively. These positive spillover effects on continuing wages take place because the reform increased the implicit wage bargaining power of these two groups of workers as a consequence of the reduction in their unemployment rates. According to our simulated results, the unemployment rates of young and older workers decreased from 34.7% and 12.6% to 33.3% and 9.8%, respectively.[15] As a result of the simulated response in the wages of new and continuing workers, the relative average wages of young and older workers increased by 3.21% and 4.22%.

The simulation also permits to separately identify and quantify the effects of each policy change. That is, we can compute the impact of changing either firing costs or payroll taxes. To calculate the impact of reducing solely payroll taxes, we simulate a scenario with no reduction in firing costs, keeping the rest of post-reform parameters constant. The results of this exercise are presented in the bottom panel of Table 7. According to our model, the direct effect of payroll taxes reduce both new hired and continuing wages because they reduce the net share of the match product obtained by the worker. That is, we should expect an increase in the wages of treated groups due to the reduction of payroll taxes during the 1997 reform. Our findings are consistent with theoretical predictions. Payroll taxes account for 68% and 63% of the increase in average wages (new entrants plus incumbents) of young and older workers with respect to middle aged workers. However, the reduction in payroll taxes only explains around 40% of the increase in the wages of new hired workers. This result takes place because, in contrast to the case of incumbent workers, lower firing costs also increase the wages of new entrants.

Notice that, although firing costs have a negative effect on incumbent's wage, when

[15] It is interesting to note that the simulated reduction in unemployment goes in line with the positive employment effects estimated by Kugler, Jimeno, and Hernanz (2002) and, more recently, by Cervini Plá, Ramos, and Silva (2010).

keeping firing costs constant in the simulation (bottom panel), the relative wage increase of incumbents is smaller than the relative wage increase that obtains from simulating the whole reform (upper panel). This result takes place because the reform only modified firing costs for new permanent contracts but not for old permanent contracts. Thus, in our simulated reform, firing costs for incumbents do not change. Moreover, since the drop in unemployment is lower when firing costs are kept constant, there is less pressure for a wage increase.

7 Final remarks

This paper provides empirical evidence of the effect on wages of two important elements of non-wage labour costs, using a labour market reform in Spain which reduced firing costs and payroll taxes after 1997 for certain population subgroups.

To gain a theoretical insight into the effects of these two provisions we extend the matching model with heterogeneous workers (Dolado, Jansen, and Jimeno (2007)) to accommodate the salient features of the reform. Since the firm does not incur in firing costs when there is no agreement on a wage between the firm and the employee in the first encounter, we permit the wage bargaining process to differ between new entrants and incumbent workers. Because of this, the model predicts a different impact of the reform on entry and continuing wages. For incumbent workers, smaller firing cost reduce wages through their decreased bargaining power of workers. Such reduction, however, benefits new entrants, for whom firing costs are not operational. The reduction of payroll taxes decreases the net share of the product obtained by the worker, which in turn increases the wages of both type of workers. In sum, the model predicts a positive impact of the reform on entry wages and an unambiguous effect for incumbent workers.

For the empirical analysis we use a unique longitudinal data set, which contains information on individual job histories from social security records and basic individual characteristics from the census. Since we have information on each and every single job spell, we avoid the possibility of aggregation bias.

Our empirical strategy exploits the substantial reduction in firing costs and payroll taxes brought about by the 1997 Spanish labour market reform for young and old workers who got a permanent job from unemployment. Since the changes did not cover all workers, we use a difference-in-differences estimator to obtain short-term causal effects. The possible sample selection bias that arises because firing cost and labour tax reductions apply only to workers transiting from unemployment to permanent employment is addressed with a two-step Heckman correction model. The first step of the model is identified with information on previous unemployment spells. Identification of the causal effects of the reform may be threatened if employers substitute workers not affected by the reform for targeted workers. We show that substitution of workers does not take place. Our estimates suggest that decreased firing costs and payroll taxes have a positive effect on wages (and also on unemployment). We find larger effects for older than for younger workers and for men than for women.

Calibrating the model and simulating the reform provides a robustness check of the estimated effects and allows to separately identify and quantify the effects of each policy change, which cannot be estimated since the two provisions changed at the same time. Simulated effects are consistent with the estimated effects, though somewhat larger in size for younger workers. Regarding the relative impact of each provision, our simulations suggest that two thirds of the increase in average wages is due to the reduction in payroll taxes, while firing costs account only for one third of the increase. However payroll taxes account for a smaller proportion of the overall wage increase for new hired workers (40%), since the reduction in firing costs also increases the wage of this group of workers.

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A Appendix

A.1 Main changes in dismissal costs and payroll taxes Due to the 1997 Reform for temporary workers

Table 8: **Principal Changes in Dismissal Cost and Payroll Tax due to the Labour Market Reform of 1997 which permit identification for Temporary Contracts**

		Dismissal cost under existing permanent contracts (pre-reform)	Dismissal cost under new permanent contracts (post-reform)	Payroll tax reductions for newly hired workers under permanent contracts after 1997
Treated group	Older (>45 years)	45 days' wages per year of seniority with a maximum of 42 months' wages	33 days' wages per year of seniority with a maximum of 24 months' wages	60% of employer contribution for 24 months, 50% thereafter
Control group	Young and Middle-aged (≤45 years)	45 days' wages per year of seniority with a maximum of 42 months' wages	33 days' wages per year of seniority with a maximum of 24 months' wages	50% of employer contribution for 24 months

A.2 Sensibility checks with wider time windows (2 years)

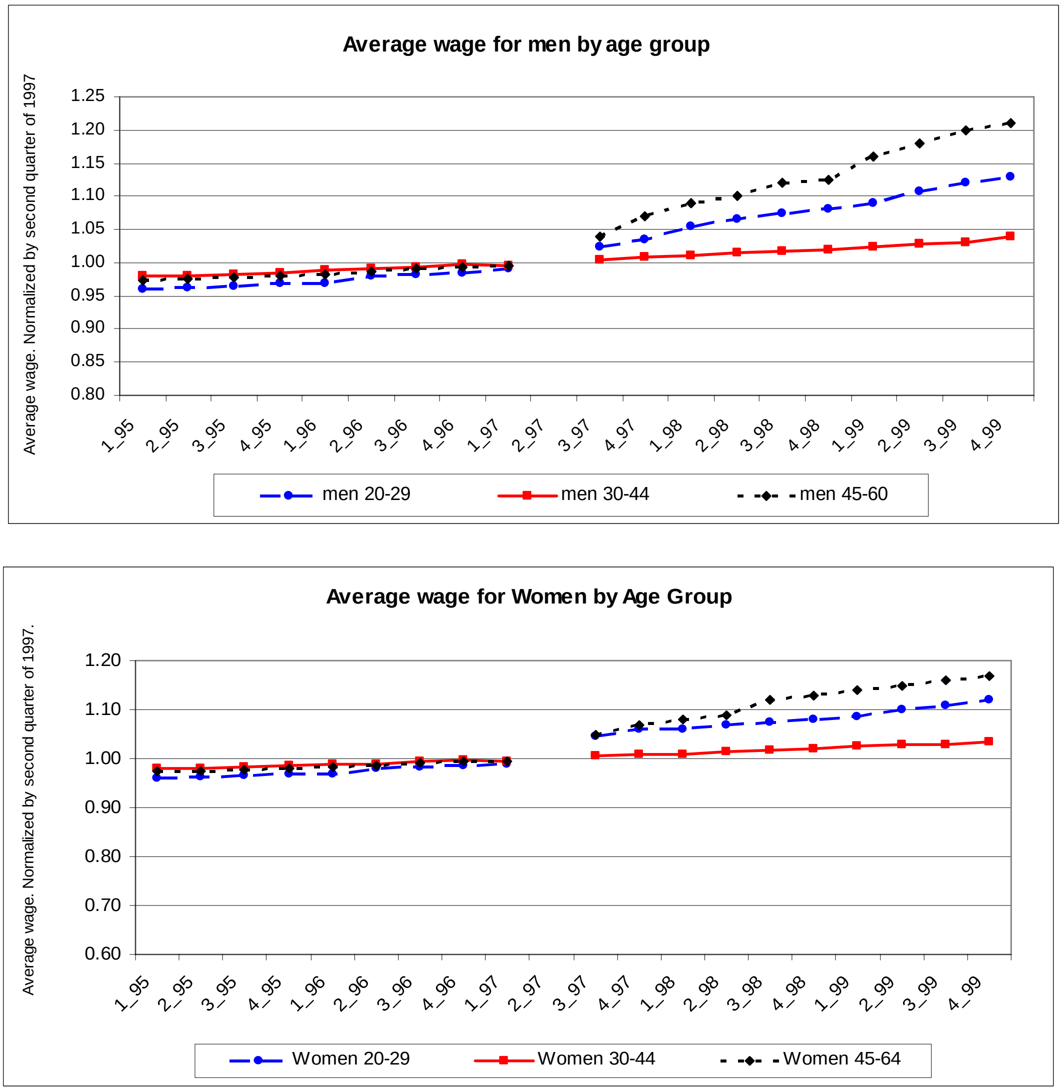
Table 9: Effects of the Reform on Wages for men and women who experience a transition from unemployment to permanent employment (**95-96 vs 98-99**)

	Men		Women	
	Coefficient	t-stat	Coefficient	t-stat
Wage equation				
Age<30	-0.253	-11.18	-0.231	-9.78
Age>45	0.122	4.63	0.042	2.49
(Age<30)*(Post 1997)	0.032	4.84	0.027	2.71
(Age>45)*(Post 1997)	0.069	4.73	0.060	3.74
Selection coeff (λ)	17.882	5.70	3.432	18.5
Selection equation				
Age<30	-0.294	-11.27	-0.245	-7.27
Age>45	-0.091	-5.99	0.028	5.85
(Age<30)*Reform	0.027	4.42	0.024	4.25
(Age>45)*Reform	0.043	13.75	0.049	7.49
Unemployment spells	-0.010	-31.23	-0.011	-19.11
Duration	-0.001	-13.53	-0.001	-21.47

Notes: All coefficients are significant at 5%. Control group are men and women aged 30 to 45 years.

A.3 Full estimates of wage and selection regressions

Figure 1: **Wage trend for treated and control groups in our sample**



Source: MCVL, own calculations.

Table 10: Effects of the Reform on Wages for men and women who experience a transition from unemployment to permanent employment

	Men		Women	
	Coefficient	t-stat	Coefficient	t-stat
Wage equation				
Age<30	-0.246	-24.94	-0.198	-15.71
Age>45	0.040	2.73	0.049	2.70
(Age<30)*(Post 1997)	0.048	2.92	0.027	3.15
(Age>45)*(Post 1997)	0.077	4.82	0.062	2.53
(a)**Education**				
Unknown	0.129	4.71	0.051	1.37
Primary incomplete	0.038	1.35	0.184	4.27
Secondary incomplete	0.225	7.42	0.414	9.78
Secondary completed	0.243	6.49	0.521	12.86
Graduate	0.561	16.48	0.712	17.86
Postgraduate and doctorate	0.663	10.16	0.898	12.69
(b)**Industry**				
Agriculture and Fish Industry.	0.525	9.34	0.373	2.81
Extractive industry	0.213	11.23	0.695	5.76
Alimentary and Drink industry	0.353	12.47	0.942	7.67
Electricity, Gas and Water industry	0.927	8.56	0.567	4.61
Trade and Sales industry	0.858	7.95	0.521	4.32
Transport industry	0.932	8.63	0.681	5.62
Finance activities	0.740	6.85	0.445	3.69
Education and sanitary	0.540	4.95	0.396	3.28
Other social activities	0.227	2.10	0.340	2.81
(c)**Occupation**				
No qualification workers	0.358	43.74	0.403	2.81
Administrative employees	0.078	5.80	0.196	18.20
Managers	0.723	59.35	0.753	49.10
Professional	0.579	46.14	0.756	34.33
Selection coeff (λ)	13.198	8.78	6.570	9.68
Selection equation				
Age<30	-0.068	-6.02	-0.111	-12.13
Age>45	-0.053	-6.09	0.035	2.55
(Age<30)*Reform	0.038	3.62	0.018	3.66
(Age>45)*Reform	0.054	13.46	0.034	8.36
Unemployment spells	-0.016	-33.85	-0.008	-15.75
Duration	-0.0001	-22.62	-0.0001	-31.15
(a)**Education**				
Unknown	0.061	0.27	0.034	2.67
Primary incomplete	0.012	0.47	0.020	0.64
Secondary incomplete	0.065	3.33	0.014	2.59
Secondary completed	0.101	4.37	0.027	3.85
Graduate	0.342	9.38	0.111	5.65
Postgraduate and doctorate	0.212	6.19	0.243	4.93
(b)**Industry**				
Agriculture and Fish Industry.	0.345	4.53	0.438	4.74
Extractive industry	0.384	5.26	0.411	5.03
Alimentary and Drink industry	0.384	5.25	0.257	3.09
Electricity, Gas and Water industry	0.062	0.86	0.089	1.07
Trade and Sales industry	0.452	0.86	0.337	4.14
Transport industry	0.272	6.21	0.131	1.60
Finance activities	0.055	3.73	0.021	2.26
Education and sanitary	0.795	2.82	0.057	2.68
Other social activities	0.295	4.03	0.200	2.44
(c)**Occupation**				
No qualification workers	0.146	25.39	0.046	6.57
Administrative employees	0.527	5.77	0.019	2.65
Managers	0.247	28.65	0.099	9.30
Professional	0.210	23.65	0.165	10.21

Notes: Control group are men and women aged 30 to 45 years.